THE MACROECONOMIC AND FINANCIAL IMPACTS OF EUROPEAN CRISIS ON SAUDI ARABIA

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Abstract

This paper examines the contagion effect caused by the European debt crisis on the Saudi Arabian stock market and the effect's spread to the real economy. Firstly, the presented analysis tests the hypothesis about the contagion effect and interdependence between the markets by means of Gregory and Hansen's co-integration in presence of structural breaks, as well as the test for common trends proposed by Stock and Watson (1988). Next, to identify the transmission channel of the European financial crisis to the Saudi Arabia real economy, we estimate a model with a dummy variable. Finally, we estimate a stylized Phillips curve to investigate the impact of the crisis on Saudi inflation rate. The empirical results indicate that Saudi Arabian stock market has been notably affected by the European crisis. The dummy variable model demonstrates that the primary transmission channel of the financial crisis to the Saudi Arabian output is international trading. In conclusion, the estimated stylized Phillips curve remains unchanged after the crisis. This finding indicates that the crisis has not affected the structure of the relationship between inflation and output.

Keywords

European Crisis Contagion; Saudi Arabia; Equity Markets; Macroeconomic Variables; Co-Integration; Structural Break

1. Introduction

The current study investigates the contagion effect of the European crisis on Saudi Arabia stock market and the effect's transmission channel to the real economy. While still in the process of forming the European Union (EU) as a unified state, major European countries have faced an unexpected economic and financial crisis. Started in Greece, the crisis has rapidly spread to the majority of the European financial and real markets. As a result of the dependence between international markets, the European crisis has a potential to affect the trading partner countries. A number of key statistics illustrate significant growth in economic operations between the EU and Saudi Arabia.

During the last decade, the import operations from the EU to Saudi Arabia have increased sharply from 13507 to 26399 billion euros. The annual growth rate, which was approximately 6.89%, also increased. Thus, the EU has become the first import

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partner. In the same period, the export operations to the EU have risen from 13165 to 28125 billion euro with the annual growth rate of approximately 7.87%.

Following the Observatory of Economic Complexity (2013)[1], the export operations of KSA are still dominated by crude oil representing 77% of the total export. Japan is the largest export market of Saudi Arabia, accounting for approximately 16% of the total merchandise export. Japan is followed by China and the United States with 15% of the total Saudi export.

The strengthening economic relationships between the EU and Saudi Arabia has motivated this investigation of the possible contagion effect of the European crisis to the Saudi financial market and the effect's subsequent transmission to the real economy.

Multiple scholar investigate the problem of contagion in financial markets and real economies, including Baig and Goldfajn (1999), Forbes and Rigobon (2002), Longin and Solnik (1995), Bekaert et al. (2005) and, more recently, Kroszner et al. (2007), Demetriou et al. (2013) and Dumontaux and Pop (2013), to name several examples.

Using a GARCH specification, Longin and Solnik (1995) demonstrate that the international market correlations are unstable during the period of 1960 to 1990. The authors also find that correlation increases in the periods of high volatility. Baig and Goldfajn (1999) also observe analogous results when they test significance of the increase in cross-market correlations during the East Asian crisis. These researchers discover that correlation in currency and sovereign spreads increases significantly during the East Asian crisis. In addition to this finding, the authors present evidence of the contagion in currency and equity markets. Forbes and Rigobon (2002) also argue that volatility has an impact on high cross-market correlations. These researchers discuss the contagion using stock market correlations during the Asian crisis of 1997, the Mexican devaluation in 1994, and the US market crash in 1987. The authors show that there is no increase in unconditional correlation coefficients during these turbulent periods. Kroszner et al. (2007) analyze the mechanisms linking financial and real activities during a financial crisis. Using data from 38 developed and developing countries that have experienced financial crises, the authors report that during a banking crisis, industries that rely heavily on external finances suffer greater losses in value added and may cause systemic defaults within Europe. Demetriou et al. (2013) perform an empirical investigation of the contagion effects on emerging equity markets during the global financial crisis and observe that links reemerged after the Lehman Brothers collapse. This finding suggests a change in investors' risk appetite. Dumontaux and Pop (2013) examine investors' reaction to the Lehman Brothers collapse in an attempt to identify a spillover effect on the surviving financial institutions.

In addition to these sources, there are numerous studies on the euro crisis's effect on the economies and financial markets of the European partners. Among these studies are works by Gerlach et al (2010), Hoque (2012), Frias (2013), Guillen (2012), De Grauwe and Yuemei (2013), De Haan (2013), Hui et al. (2013), Ahmad et al. (2013), Mallick

[1] http://atlas.media.mit.edu/country/sau/

and Sousa (2013), Sensoy et al (2014), Garcia and Ramos (2015), Shen et al. (2015) and Tola and Wälti (2015).

Gerlach et al. (2010) study the determinants of sovereign bond spreads in the euro area since the introduction of the euro currency. The authors show that for several peripheral European countries, such as Greece, Ireland and Portugal, a jump in their sovereign CDS spreads has transmitted to their banks' CDS spreads. Hoque (2012) studies the effects of the European crisis on currency markets. More specifically, this researcher implements a classic co-integration test to identify relationship between currencies. Guillen (2012) analyzes the European crisis and its connection to the global economic and financial crisis. Musialkowska et al. (2012) discuss the functioning of the Single Market in crisis conditions and the actions taken by the EU to alleviate the negative effects of the crisis.[2] More recently, De Grauwe and Yuemei (2013) discuss the theory of fragility of the Eurozone and propose to test this theory using a multiple linear regression. Mink and De Haan (2013) also utilize a linear regression to identify the twenty days with extreme returns on Greek sovereign bonds and categorize the news events during those days into news concerning Greece and news regarding the prospects of a Greek bailout. Hui et al. (2013) define a measure of a spillover effect of option-implied correlation between indices during the European debt crisis. The authors implement an analytic method and find an interdependence among various markets during the crisis. Ahmad et al. (2013) examine the financial contagion in an emerging market setting by investigating the contagion effects of GIPSI (Greece, Ireland, Portugal, Spain and Italy), USA, the UK and Japan markets on BRIICKS (Brazil, Russia, India, Indonesia, China, South Korea and South Africa) stock markets. During the Eurozone crisis period (October 19, 2009–January 31, 2012), the empirical results indicate that among all GIPSI countries, Ireland, Italy and Spain appear to be more contagious for BRIICKS markets than Greece. Mallick and Sousa (2013) examine the real effects of the financial stress in the Eurozone. The authors use a Bayesian Structural VAR and a Structural VAR (SVAR) and determine that unexpected variation in financial stress conditions plays an important role in explaining output fluctuations and therefore demands an aggressive response from the monetary authority to stabilize output. Sensoy et al. (2014) use a dynamic approach to investigate a relationship between Turkey and the European countries during the global financial crisis. These researchers' results demonstrate that Turkey is not immune to the global financial conditions. Moreover, the authors discover a significant integration between Turkey and the major European economies in terms of risk perception after the global financial crisis. Shen et al. (2015) employ a Kalman filter approach to estimate the time-varying correlation coefficients in their examination of the contagion effect on China's stock market during the European debt crisis. Their results indicate that after introducing the control over the macro fundamental variables and global shocks, the contagion effect of the crisis on investors' psychology in the Chinese capital market was limited. Tola and Wälti (2015) combine the standard contagion test by Favero and Giavazzi (2002) with a narrative approach to test for the presence of the financial

[2] For an in-depth analysis of the European crisis and its main causes, see Blankenburg (2013) and Aizenman et al (2013).

contagion during the crisis in the euro area. The authors attempt to separate global and the euro area shocks from country-specific shocks. Their results demonstrate that three quarters of the country-specific shocks are contagious over the entire sample period. However, the fraction of the contagious country-specific shocks has fallen markedly after the “whatever it takes” announcement by ECB President Draghi in July 2012 that proposed fighting unreasonably high government borrowing costs.

The most prominent of the above-mentioned empirical studies are based on the definition of the contagion proposed by Forbes and Rigobon (2002). These researchers define the contagion as an increase in correlation during periods of crisis. Nonetheless, the analysis in their paper is consistent with the definition of the contagion proposed by Dungey et al. (2005) and discussed in Stracca (2014) and Tola and Wälti (2015). In these works, the crisis is a transmission of unanticipated local shocks to another country or market.

The definition of the contagion has a direct influence on the employed approach, particularly in the case of measuring the contagion between markets. Those studies that use the definition proposed by Forbes and Rigobon (2002) tend to implement models that measure coefficients of correlation or estimate dependence parameters using a Dynamic Conditional Correlation (DCC) multivariate GARCH specification. For example, one can refer to the works by Bekaert and Harvey (2000), Boyer et al. (2006), Chiang et al. (2007), Lucey and Voronkova (2008), Jeon and Moffett (2010), and Manolis et al. (2011).

Although the multivariate GARCH models are able to measure conditional correlations and thus the contagion effect, they have an underlying assumption that innovations follow a symmetric multivariate normal or Student’s t-test distribution. Therefore, these models fail to reflect the financial markets with the asymmetric tail dependence (Wang et al., 2013). Moreover, when a number of multivariate GARCH specifications are used, each of them might suggest a different pattern of correlation dynamics (Kroner and Ng, 1998; Bauwens et al., 2006). Hence, the results from such models can be highly misleading. Additionally, dynamic conditional correlation models tend to exhibit highly unstable conditional correlation patterns and erratic behavior (Füss and Glück, 2012). Another competing approach commonly used to investigate the contagion between markets is the vector autoregressive model (VAR) (Bekaert et al., 2005; Longin and Solnik, 2001;Blatt et al.,2015).

Other techniques can also be applied to measure the contagion between markets. For example, Dungey et al. (2010) and Baele and Inghelbrecht (2010) use a factor model, while Rodriguez (2007) and Aloui et al. (2011) use a copula approach.

This paper utilizes a slightly different methodology from those proposed in the previous literature. First, we employ the Gregory and Hansen co-integration to identify a long-run relationship among European and Saudi Arabian economic and financial variables. This co-integration approach allows the detection of co-integration with the presence of a structural break. Second, to investigate the dependence between the Saudi and European financial markets during the crisis period, we apply the test for common trend proposed by Stock and Watson (1988). Third, we estimate a model with a dummy variable to identify the transmission channel of the financial crisis to the real economy. Final, we estimate a stylized Phillips curve to examine the impact of the crisis on inflation.

To the best of our knowledge, the literature that investigates the contagion of the European crisis to the Saudi Arabia market is extremely limited. For example, Moosa (2010) uses a co-integration test to study the stock market contagion from the United States to the markets of the GCC countries during the period of 2007 to 2008. More specifically, this researcher notes that the Saudi Arabia financial market is affected by the decline, but not by the increase, of the American index (Dow Jones). This observation suggests that the effect of volatility of the American index on the Saudi market is asymmetric.

Suliman (2011) applies the GARCH method to examine the evidence of the contagion in the GCC countries from 1960 to 2002. This researcher finds that Saudi Arabia first suffered from the contagion effect of the US stock market crash in 1987 and in 1997.

Neaime (2012) studies the contagion of the global financial crisis for the case of the MENA countries, including Saudi Arabia. Essentially, he shows that the Saudi Arabia financial market has had the lowest correlation with other financial markets during the crisis period.

Despite these works, we emphasize that there exists no systematic empirical research addressing the impact of the financial crisis in the European countries on Saudi Arabia.

This study is organized as follows. In methodology section, we present our approach to examining the impact of the European Crisis on the Saudi Arabian economy. Mainly, we focus on the long run relationship between the two markets. The analysis section describes the data and presents our preliminary analysis. The results section discusses the empirical results. Finally, we conclude the study with the primary findings of the research and possible policy implications.

2. Methodology

To inspect the impact of the European crisis on both financial and macroeconomic levels, we utilize various modeling strategies. First, we apply unit root tests to study the stationary of the financial and economic series. Specifically, we use approach by Zivot and Andrews (1992) to detect breaking dates. Second, we perform a co-integration test in presence of structural break to study the long-run relationship between the European and the Saudi stock markets. In this study, we also propose a vector autoregressive (VAR) model estimation using a structural analysis, such as impulse responses and forecast error variance decompositions, to examine the contribution of each European index in the variance of TASI after, during, and before the European crisis period. Last, we study the dependence between the markets in the crisis period with the test for a common trend proposed by Stock and Watson (1988).

In particular, , we describe in this section the co-integration test in presence of structural break. We also present the common trend test suggested by Stock and Watson (1988).

2.1. Co-integration test in presence of structural break (financial crisis)

Co-integration, as defined and developed by Granger (1981) and Engle and Granger (1987), is a property of some nonstationary time series. If two or more nonstationary time series are co-integrated, a linear combination relationship being stationary is said to exist. Co-integration can be also defined as a mathematical formulation of the long run relation between economic variables, especially the presence of the stationary long-run relations even if the variables themselves are nonstationnary.

Tests of co-integration were developed for the first order co-integrated series I(1) by Shin (1994), based on the stationary test in Kwiatkowski, Phillips, Schmidt and Shin (1992), as well as by Saikkonen and Luukkonen (1993), Xiao and Phillips (2002), and others. Specifically, Gregory and Hansen (1996) developed residual-based tests for co-integration in models with regime shifts. These tests seem to be suitable to our empirical study in view of the presence of a structural break in our data series due to the crisis and empirically proved by Zivot and Andrews's unit root test.

We start first introducing the Gregory and Hansen's methodology which allow for co-integration with structural break. They have developed four single-equation regression models for the observed data $y_t = \left(y_{1t}, y_{2t}\right)$ where y_{1t} is real-valued and y_{2t} is an *m*-vector. The standard model of co-integration can be written as follow:

$$y_{1t} = \mu + \beta_t + \alpha^T y_{2t} + e_t, \quad t = 1,...,n \tag{1}$$

Where y_{2t} is I(1) and e_t is I(0).

In this equation the parameters μ and α describe the *m*-dimensional hyperplane towards which y_t tends over time. In this modeling these two parameters are assumed to be time invariant to detect a long run relationship. However, in many time series, the presence of the structural break affect the parameters μ and α , then the co-integration is captured for some periods of time and shift to a new long run relationship after the break date.

To introduce the structural break in co-integration test, it is appropriate to specify the dummy variable:

$$\varphi_{t\tau} = \begin{cases} 0, & if \quad t \le [n\tau] \\ 1, & if \quad t > [n\tau] \end{cases}$$

Where the unrevealed parameter $\tau \in (0,1)$ indicates the timing of the breakpoint.

In this context, Gregory and Hansen (1996) discussed three models with presence of structural break;

Model 1: Level shift $$y_{1t} = \mu_1 + \mu_2\varphi_{t\tau} + \alpha^T y_{2t} + e_t, \quad t = 1,...,n \tag{2}$$

This model indicates a presence of a level shift in the co-integration relationship; this can be explained by a change in the intercept μ while the slope remains unchanged.

Model 2: Level shift with trend $$y_{1t} = \mu_1 + \mu_2\varphi_{t\tau} + \beta t + \alpha^T y_{2t} + e_t, \quad t = 1,...,n \tag{3}$$

In this specification μ_1 represents the intercept before the breakpoint and μ_2 indicates the change in the intercept over the shift.

Model 3: Regime shift $$y_{1t} = \mu_1 + \mu_2\varphi_{t\tau} + \alpha_1^T y_{2t} + \alpha_2^T y_{2t}\varphi_{t\tau} + e_t, \quad t = 1,...,n \tag{4}$$

In this specification α_1 expresses the co-integrating slop coefficients before the regime shift and α_2 denotes the change in the slope coefficients.

The Gregory and Hansen (1996) test consists of null hypotheses of no co-integration against the alternatives in the three last models presented above. The methodology proposed to treat regime shift is inspired from Banerjee et al. (1992) and Zivot and Andrews (1992). The co-integration test statistic is computed for each possible regime shift, and takes the smallest value across all possible breakpoints. The analysis of co-integration stand on the three test statistics; $Z_\alpha^* = \inf_{\tau \in T} Z_\alpha(\tau)$, $Z_t^* = \inf_{\tau \in T} Z_t(\tau)$, and $ADF^* = \inf_{\tau \in T} ADF(\tau)$, where $Z_\alpha(\tau)$ and $Z_t(\tau)$ correspond to the Phillips test statistics (for more details see Phillips, 1987).

2.2 Testing for Common trends: Stock and Watson (1988)

To investigate the dependence between the Saudi and European financial markets, in crisis period, we apply the test for common trend proposed by Stock and Watson (1988). The approach can be presented as follows.

Let X_t denote a $n \times 1$ time series variable that is co-integrated of order $(1,1)$. That is, each element of X_t, is integrated, but there are r linear combinations of X_t, that are stationary. The change in X_t, is assumed to have the co-integrated vector moving average representation:

$$\Delta X_t = \mu + C(L)\varepsilon_t, \quad \sum_{j=1}^{\infty} j\left|C_j\right| \prec \infty, \tag{5}$$

Where $C(z) = \sum_{i=0}^{\infty} C_i z^i$ with $C(0) = I_n$, ε_t is iid with mean 0 and covariance matrix G , L is the lag operator and $\Delta \equiv 1 - L$. $C(1)$ is assumed to have rank $k \prec n$, so X_t is cointegrated; that is, there is an $n \times r$ matrix α (where $r = n - k$) such that $\alpha' C(1) = 0$ and $\alpha' \mu = 0$.

A representation for the stationary linear combinations $\alpha' X_t$ is readily obtained from (5). Let $v_t = G^{-1/2}\varepsilon_t$ and $\xi_t = \sum_{s=1}^{t} v_s$, approve the usual statement (e.g., Dickey and Fuller 1979) that $\varepsilon_t = 0$ $(s \leq 0)$, and allow X_t to have a nonrandom initial value X_0.

Then recursive substitution of (5) gives:

$$X_t = X_0 + \mu t + C(1)G^{1/2}\xi_t + C^*(L)G^{1/2}v_t$$

(6)

Where $C^*(L) = (1-L)^{-1}(C(L) - C(1))$ so that $C_j^* = -\sum_{i=j+1}^{\infty} C_i$. Because $\alpha' C(1) = 0$ and $\alpha' \mu = 0$, it follows that:

$$Z_t \equiv \alpha' X_t = \alpha' X_0 + \alpha' C^*(L) G^{1/2} \nu_t \qquad (7)$$,

The cointegrated process X_t has an alternative representation in terms of reduced number of common random walks plus a stationary component. This "common trends" representation is readily derived from (6). Because $C(1)$ has rank $k \prec n$, there is an $n \times r$ matrix H_1 with rank r such that $C(1)H_1 = 0$. Furthermore, if H_2 is an $n \times k$ matrix with rank k and columns orthogonal to the columns of H_1, then $A \equiv C(1)H_2$ has rank k. The $n \times n$ matrix $H = (H_1 H_2)$ is nonsingular and $C(1)H = (0 \;\; A) = AS_k$, where S_k is the $k \times n$ selection matrix $\lfloor 0_{k \times (n-k)} I_k \rfloor$, where $0_{k \times (n-k)}$ is a $k \times (n-k)$ matrix of zeros. In addition, because $\alpha' C(1) = 0$ and $\alpha' \mu = 0$, μ lies in the column space of $C(1)$ and can be written $\mu = C(1)\tilde{\mu}$, where $\tilde{\mu}$ is an $n \times 1$ vector. Thus (7) yields the common trend representation for X_t:

$$\begin{aligned} X_t &= X_0 + C(1)\left[\tilde{\mu} t + G^{1/2}\xi_t\right] + C^*(L) G^{1/2} \nu_t \\ &= X_0 + C(1)H\left[H^{-1}\tilde{\mu} t + H^{-1} G^{1/2} \xi_t\right] + a_t \\ &= X_0 + A\tau_t + a_t, \quad \tau_t = \pi + \tau_{t-1} + \upsilon_t, \qquad (8) \end{aligned}$$

Where $a_t = C^*(L) G^{1/2} \nu_t$, $\tau_t = S_k H^{-1} \tilde{\mu} t + S_k H^{-1} G^{1/2} \xi_t$, $\pi = S_k H^{-1} \tilde{\mu}$

The common trends representation expresses X_t as a linear combination of k random walks with drift π, plus some transitory components, a_t, that are integrated of order 0.

A usual approach to test k versus m common stochastic trends would be to examine the first-order serial correlation matrix of X_t.

Stock and Watson propose to examine functions of regression statistics of a linear transformation of X_t, denoted by Y_t, chosen so that under the null hypothesis the first $n-k$ elements are not integrated, whereas the final k elements of Y_t can be expressed in terms of the k separate trends. Specifically, let $Y_t = DX_t$, where $D = \left[\alpha \alpha^+\right]'$, where α^+ is an $n \times k$ matrix of constants selected so that $\alpha^{+'} \alpha = 0$ and $\alpha^+ \alpha^{+'} = I_k$. The first $n-k$ elements of Y_t correspond to Z_t in (7). Let W_t denote the final k integrated elements of Y_t. It follows from (5) that

$$\Delta W_t = \alpha^{+'} \mu + u_t, \qquad (9)$$

where $u_t = \overline{C}(L)\nu_t$, with $\overline{C}(L) = \alpha^{+'} C(L) G^{1/2}$. Combining (6) and (9),

$$\Delta_k Y_t = \delta + F(L)\nu_t, \qquad (10)$$

Where $\Delta_k = \begin{bmatrix} I_{n-k} & 0 \\ 0 & \Delta I_k \end{bmatrix}$, $\delta = \begin{bmatrix} \alpha' X_0 \\ \alpha^{+'} \mu \end{bmatrix}$, $F(L) = \begin{bmatrix} \alpha' C^*(L) G^{1/2} \\ \overline{C}(L) \end{bmatrix}$.

Recursive substitution of (6) shows that Y_t can be represented as;

$$Y_t = \begin{bmatrix} \alpha' X_0 \\ \alpha^{+'} X_0 \end{bmatrix} + \begin{bmatrix} 0_{k\times(n-k)} \\ \alpha^{+'} \mu \end{bmatrix} t + \begin{bmatrix} 0_{k\times(n-k)} \\ \overline{C}(1) \end{bmatrix} \xi_t + \begin{bmatrix} \alpha' C^*(L) G^{1/2} \\ \overline{C}^*(L) \end{bmatrix} \nu_t \tag{11}$$

$$= \beta_1 + \beta_2 t + \beta_3 \xi_t + \beta_4(L) \nu_t$$

Where $\overline{C}^*(L) = (1-L)^{-1}\left(\overline{C}(L) - \overline{C}(1)\right)$

In terms of W_t, a test of k versus m common trends becomes a test of whether $C(1)$ has rank k against the alternative, that it has rank m. To motivate the proposed tests, suppose that $X_0 = \mu = 0$, and consider the result of regressing W_t onto W_{t-1}. Under the null hypothesis, W_t is a linear combination of W_t integrated processes, so Φ, the probability limit of

$$\tilde{\Phi} = \left[\sum W_t W'_{t-1}\right]\left[\sum W_{t-1} W'_{t-1}\right]^{-1} \tag{12}$$

has k real unit roots. Under the alternative W_t includes m integrated variables and $k-m$ nonintegrated variables, or equivalently W_t has $k-m$ linearly independent cointegrating vectors. Thus under the alternative Φ has only m unit eigenvalues corresponding to the m integrated variables, and $k-m$ eigenvalues with modulus (and therefore with real parts) less than 1. Letting λ_{m+1} denote the eigenvalue of Φ with the $(m+1)^{\text{th}}$ largest real part, our null and alternative hypotheses are $H_0 : real(\lambda_{m+1}) = 1$ versus $H_1 : real(\lambda_{m+1}) \prec 1$.

Stock and Watson proposed two statistics to test these hypotheses. The first based on filtering the data and the second is based on correcting the OLS autoregressive matrix $\tilde{\Phi}$.

For more details about these tests one can refers to Stock and Watson (1988).

3. Data and Preliminary Analysis

3.1. Data

In our study, we use the primary datasets that cover the period from January 2005 to September 2013 and can be retrieved from Bloomberg. This period includes the most recent major European crisis. The data render historical daily market indices of the Saudi Arabia and European financial markets (France, Germany and Italy as the main trading partners of Saudi Arabia and Greece, Ireland, Portugal and, Spain as the countries that are the most affected by the crisis). Additionally, the data cover quarterly macroeconomic series, specifically, GDP, importation, exportation volumes, and inflation rates.

The sample data were divided into the three subsets: the first one is before European crisis (January 2005 to June 2007), the second one is during the European crisis period (July 2007 to June 2008), and the last one is after the subprime crisis (July 2008 to 2013). The definition of these different periods has been theoretically based on Blackburn's (2008) and Kolb's (2011) decomposition and empirically examined by Zivot and Andrews' test.

3.2. Preliminary analysis

A preliminary analysis shows the evolution of different variables. The financial series plots in Fig. 1 illustrate that the Saudi index and the indices of the European countries (after a logarithmic transformation) evolve together over time. Fig. 1 also shows the presence of a breakpoint between 2008 and 2009, which can be interpreted as the crisis period. In this particular time period, several European countries were suffering the worst economic crisis, regarding the rate of growth of real GDP per head, since the period between the Great Depression of 1929 to 1930 until 1946. Of course the negative consequences of this diminution, on population welfare, were then worse than now, because in the period 2008-2016 the level of real GDP per head is more than five times higher than the values of year 1929 in many of the major European economies.

This breakpoint is also graphically shown for the Saudi Arabia index.

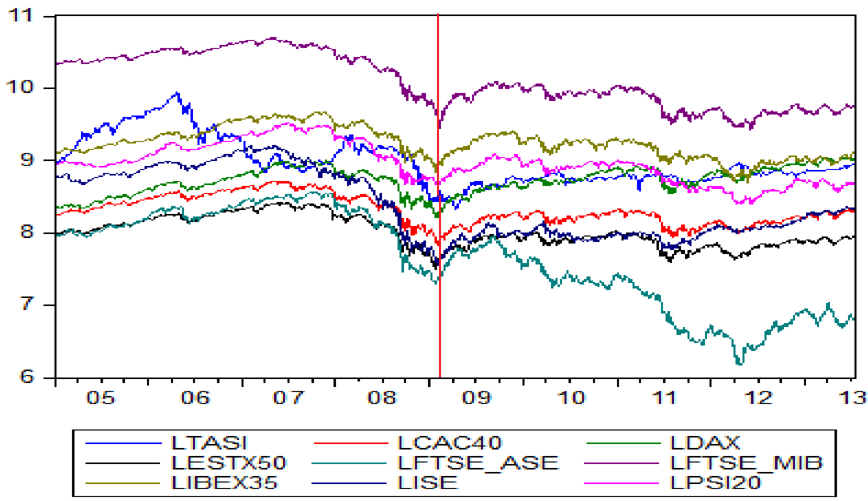

Fig. 1. Daily historical price indices of Saudi Arabia and the European countries

Regarding the economic variables, we particularly analyze the evolution of Saudi Arabia output, such as the crude oil price and international trade. As shown in Fig. 2, the spike in oil prices by approximately 8.5% in spring 2008 was associated with a surge in oil production, as well as by the high real GDP growth rates with average values over 4.5%. This trend reversed in September and December of the same year when the oil prices and GDP decreased to 13.34% and 23.15, respectively. During a six month period (more specifically, 2008: Q4, and 2009:Q1), the GDP fell 32%. The low oil price persisted throughout the year of 2009; next, it started to increase to approximately $108.65 and stabilized until 2013. Another observation is that the real GDP has almost the same trend or behavior as that of the crude oil price.

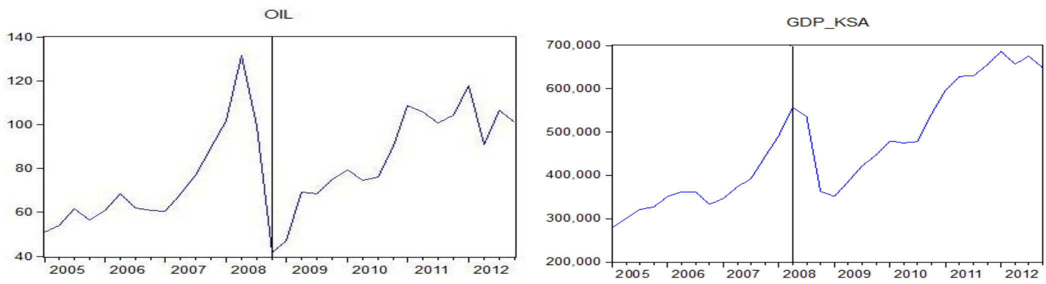

Fig. 2. The behavior of oil prices and Saudi Arabia GDP between 2005 and 2013

Historically, the US has been the principal export partner of Saudi Arabia. However, since the year 2000, the commercial ties between Saudi Arabia and other countries have experienced a significant transformation. Specifically, Saudi Arabia's has developed stronger relationships with China, India, Brazil and, particularly, the European countries. For example, Germany has become the third importer partner to the Kingdom of Saudi Arabia (KSA) with 9.8% of the total export, Italy and France share 4.7% of the total export and Spain has 2%.

Fig. 3 illustrates that Germany is the first exporter to Saudi Arabia followed by France and later Italy. These three countries are considered the primary European importers. We also observe that the import volume from the countries which are the most affected by the crisis (e.g., Greece, Ireland and Portugal) has become limited.

However, export to the European countries is still relatively insignificant in comparison to that to other developed countries. As shown in Fig. 3, the first Saudi export destination in Europe is Italy with 3% of the total export followed by France, representing 2.7% of the total export. We also note that export operations to the affected countries and Spain are, 2% and Greece 0.17%, respectively.

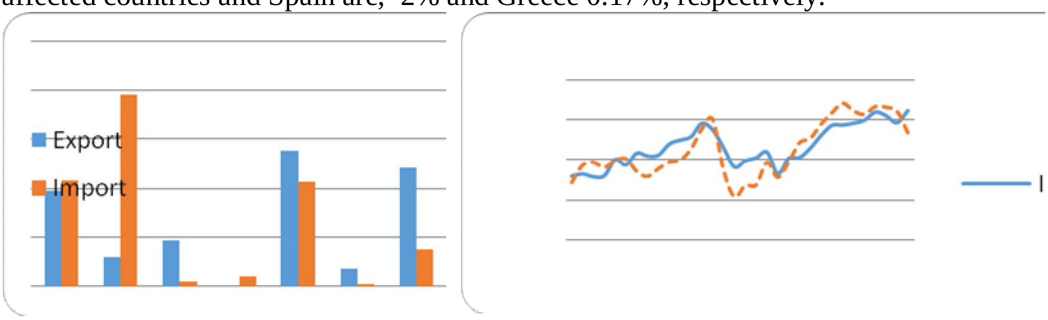

Fig. 3. The evolution of export and import operations between the European countries and Saudi Arabia

To examine the evolution of inflation, we observe the evolution of the average inflation rate as measured by annual changes in the cost of living indices before, during, and after the European crisis period. As shown in Fig. 4, before the European crisis (2005-2008), inflation rose rapidly, despite the dynamic monetary tightening started by the Saudi Arabia monetary agency. The measure included sharp expansion in commercial bank reserve requirements and boosting the issuance of treasury bills. In that period, the principal causes of the high inflation in Saudi Arabia were food prices and housing[3]. In fact, the increase in food prices was primary due to the rise in domestic demand for the commodity from the exporting countries as well as droughts.

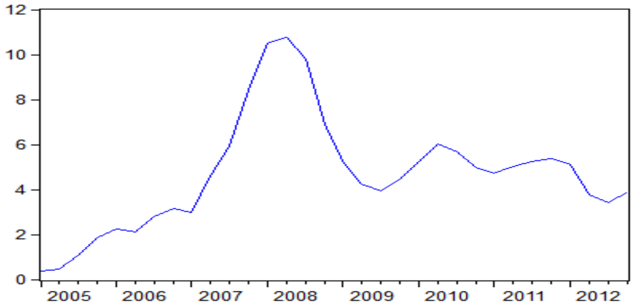

Fig. 4. Saudi Arabia: Consumer Prices (Log change)

The same Fig. 4 shows an inflation peak at 11.1% in July 2008 and inflationary pressure decline in the second half of 2008. This particular period corresponds to the European crisis. As mentioned in the Saudi Arabia Report 2009, the inflation level of the fourth quarter of 2008 can be viewed as atypical. The reduction in price pressures could be caused by the positive effect of the European crisis on import due to weakness of the euro currency. During 2009, the inflation increases again with an average rate of 5.1%. Additionally, most inflation rates by major groups exceeded their average levels recorded in the last five years.

The preliminary analysis of different variables suggests the presence of a breakpoint that can be detected using certain empirical tests. To achieve this goal, we use various unit root tests to study the stationary of time series. Particularly, we focus on the test by Zivot and Andrews (1992), the results of which enable us to detect the breaking dates. Next, we investigate the impact of the European crisis on financial and economic variables empirically.

4. Empirical Results

4.1. Timing of the crisis

Before we proceed to the investigation of the impact of the European crisis on both Saudi Arabia financial and economic levels, it is instrumental to detect empirically the timing of the European crisis in each country and, particularly, in the Kingdom. To this end, first we apply Augmented Dickey Fuller (1979) and Philips Perron (1988) tests on natural logarithm for each time series. These unit root tests examine whether the financial and macroeconomic variables are stationary or moving with time. Second, we apply the test by Zivot and Andrews (1992) to search for a breakpoint endogenously and test for the presence of a unit root if the process has a broken constant or trend. We chose this test because it is robust and more powerful than the ADF and PP tests. The results for the latter are reported in Table 1 and Table 2, respectively.

Table 1. Results of ADF and PP test on financial and economic series.

[3] The Report: Saudi Arabia 2009, Oxford Business Group retrieved from http://books.google.com.sa/books/about/ The_Report_Saudi_Arabia_2009.html?id=v6s3wtsA-HgC&redir_esc=y, at 15 October 2013.

Series	ADF level	PP level	ADF first diff	PP first diff
Financial variables				
EURO STOXX 50	-2.36 (0.400)	-2.021 (0.588)	-49.644 (0.000)	-49.967 (0.000)
TASI	-1.817 (0.696)	-1.753 (0.726)	-42.064 (0.000)	-42.218 (0.000)
CAC 40	-2.140 (0.522)	-2.040 (0.578)	-50.085 (0.000)	-50.519 (0.000)
DAX	-2.007 (0.596)	-2.085 (0.553)	-46.843 (0.000)	-47.07 (0.000)
FTSE MIB	-1.993 (0.428)	-2.209 (0.483)	-47.248 (0.000)	-47.334 (0.000)
FTSE/ASE	-2.234 (0.469)	-2.217 (0.479)	-44.258 (0.000)	-44.210 (0.000)
ISE	-0.665 (0.974)	-0.582 (0.979)	-45.448 (0.000)	-45.480 (0.000)
PSI-20	-2.048 (0.573)	-2.060 (0.567)	-47.475 (0.000)	-44.404 (0.000)
IBEX 35	-2.598 (0.281)	-2.457 (0.349)	-46.265 (0.000)	-46.472 (0.000)
Economic variables				
GDP_KSA	-1.317 (0.60)	-0.916 (0.769)	-3.932 (0.0053)	-3.957 (0.0049)
IMP_Europ	-1.485 (0.527)	-6.052 (0.000)	-1.449 (0.545)	-6.066 (0.000)
Exp_Europ	-1.932 (0.314)	-4.732 (0.0007)	-2.072 (0.256)	-5.028 (0.0003)
Inf_KSA	-4.260 (0.0023)	-	4.074 -(0.0036)	-

Levels and Diff denote the Augmented Dickey-Fuller and Phillips-Perron t-tests for a unit root in levels and first differences respectively. Number of lags was selected using the AIC criterion. Where terms in parentheses, denote p-values

Table 2. Results of Zivot and Andrews test on financial series.

Series	Minimum *t*-stat	Shift dates
EURO STOXX 50	-3.962	05/16/2008
TASI	-7.023	03/14/2007
CAC 40	-4.137	05/28/2008
DAX	-4.140	05/06/2008
FTSE MIB	-3.654	05/30/2008
FTSE/ASE	-2.725	10/22/2009
ISE	-4.455	05/27/2008
PSI-20	-2.844	05/28/2008
IBEX 35	-3.076	05/23/2008

Note: For each time series the minimum *t*-statistics, the equations (2), (3) or (4) was estimated with the breakpoint, ranging from $t = 2$ to $t = T - 1$. The minimum *t*-statistic reported is the minimum over all $T - 2$ regressions. The shift dates indicate the presence of structural break in each studied series. The critical values for *t*-statistic detecting a break are -5.34, -4.8, and -4.58 at 0.01, 0.05 and 0.1 level respectively.

Table 3. Results of Zivot and Andrews test on macroeconomic series.

Series	Minimum *t*-stat	Shift dates	Series	Minimum *t*-stat	Shift dates
Import France	-4.786**	Q4-2008	Export France	-6.046***	Q4-2008
Import Germany	-2.252	Q4-2008	Export Germany	-3.842	Q2-2009
Import Greece	-4.598*	Q4-2008	Export Greece	-5.347***	Q4-2008
Import Ireland	-4.809**	Q1-2010	Export Ireland	-4.393	Q3-2009
Import Italy	-3.727	Q3-2008	Export Italy	-4.592*	Q3-2008
Import Portugal	-3.565	Q4-2008	Export Portugal	-3.760	Q1-2011
Import Spain	-3.375	Q4-2008	Export Spain	-5.463***	Q3-2008
Aggregate Import	-4.673**	Q4-2008	Aggregate Export	-4.621*	Q4-2008
GDP	-5.769***	Q4-2008	Inflation	-5.228**	Q1-2007
Interest rate	-2.8322	Q1-2010	-	-	-

Note: For each time series the minimum *t*-statistics, the equations (2),(3) or (4) was estimated with the breakpoint, ranging from $t = 2$ to $t = T - 1$. The minimum *t*-statistic reported is the minimum over all $T - 2$ regressions. The shift dates indicate the presence of structural break in each studied series. The critical values for *t*-statistic detecting a break are -5.34, -4.8, and -4.58 at 0.01, 0.05 and 0.1 level respectively.

The results of ADF and Phillips and Perron tests are confirmed by the Zivot and Andrews test. The test also shows that the null hypothesis of a unit root cannot be rejected, while a unit root is rejected for the first difference of the series. Thus, the series can be regarded as realizations of stochastic I(1) variables. These results are very important because for the majority of financial series, the breakpoint is detected in May of 2008. Our results are in agreement with those obtained by Kolb (2011) and Thao and Daly (2012). Particularly, the breakpoint date detected on the Saudi index is in October of 2009.

Additionally, the Zivot and Andrews test results provided in Table 3 show that the null hypothesis of a unit root cannot be rejected for the Saudi macroeconomic series, while a unit root is rejected for the first difference of the series. Therefore, the series can be regarded as realizations of stochastic I(1) variables. The results also demonstrate the presence of a structural break for the output and import export operations in Saudi Arabia. The breakpoint date is mostly observed in the fourth quarter of 2008, which enables us to deduce the presence of financial transmission to the real economic area.

In contrast, for the interest rate variable, the Zivot and Andrews test shows that the breakpoint detected in Q1 of 2010 is statistically insignificant. Hence, the European crisis does not directly affect interest rate in Saudi Arabia. This fact is primarily attributable to the mechanism of fixing of the interest rate adopted by the Saudi Arabia Monetary Agency (SAMA). SAMA has been proactive in mitigating the impact of the crisis on domestic system liquidity. As the financial contagion started to spread to the Saudi market, SAMA lowered the reserve requirement, cut its interest rates (repo rates) and supplied liquidity to the banks through direct and fiduciary deposit placements. The Saudi interest rate has been reduced from 5.5% to 2% and the reverse repo rate from 4% to 0.25% since October 2008. These measures have predominantly helped public confidence in the banking system. We note that the interest rate policy of SAMA is strongly related to the decisions of the American central bank.

Although identification of the breakpoint is important, it is insufficient to make a conclusion regarding the crisis transmission to the economic variables and, particularly, on Saudi Arabia output. For this reason, we complement these results with those obtained from the estimated models.

4.2. Financial impact of the Eurozone crisis

To examine the reason for the gap between the timing of the European crisis and the breakpoint that happened in Saudi Arabia financial index, the following important hypotheses can be proposed and discussed:

H_1. The European crisis does not affect the Saudi financial market immediately. The reaction might manifest one year after the crisis.

H_2. The breakpoint is endogenous to the Saudi financial market.

To test the first hypothesis, we propose a co-integration test with the presence of structural break. Moreover, we propose to study the variance decomposition in three periods (before, during, and after the crisis) to examine the extent of the crisis impact on the Saudi index.

4.2.1. Co-integration with the presence of structural break

A long-run relationship between two stock market indices, i and j, can be written as follows:

$$\ln(P_t^i) = \beta_0 + \beta_1 \ln(P_t^j) + \varepsilon_t .$$

If the prices of two indices after a logarithmic transformation are co-integrated, the error term ε_t is stationary. In this case, we conclude that there exists a long run relationship among the two series. Table 4, in the Annex, shows the Gregory and Hansen co-integration test results after the log transformation of the variables.

The co-integration test results show that the long-run relationship between the indices after a logarithmic transformation is rejected for all the specifications. This result is not surprising because the variables $\ln(P_t^i)$ and $\ln(P_t^j)$ are both non-stationary. This feature is observed for the majority of the financial time series. For this reason, we proceed by a logarithmic differentiation for the time series to obtain the integrated series of the first order. Next, we apply the Gregory and Hansen test to study the long run relationship. The results are demonstrated in Tables 5 and 6, in the Annex.

Table 5 provides the results of the Gregory and Hansen test applied to the logarithm differentiated series. In this test, we consider the Saudi index (TASI) as dependent, and in Table 6, it is presented as independent.

The results show that the co-integration between the time series is accepted simultaneously with the presence of structural break obtained by the Zivot and Andrew test. Moreover, this co-integration is bi-directional. Next, we support a long-run relationship among the Saudi index and the European indices. This link is particularly obvious in the crisis period indicated by the point break. The co-integration results with the presence of structural break reveal a distinct impact of the European crisis on the Saudi financial market. To illustrate it even more vividly, we propose to investigate the extent of the crisis using an impulse response analysis and a variance decomposition of the Saudi index. The variance decomposition is compared among the three periods (after, during, and before the crisis).

4.2.2. Impulse response analysis

The impulse response function is used as a tool to estimate the degree and the timing of the effect of changes in the European indices or shocks on the Saudi index. Particularly, we analyze the effect of changes in CAC40, DAX and Euro stocks50 on TASI. The impulse response plots are presented in Fig. 5.

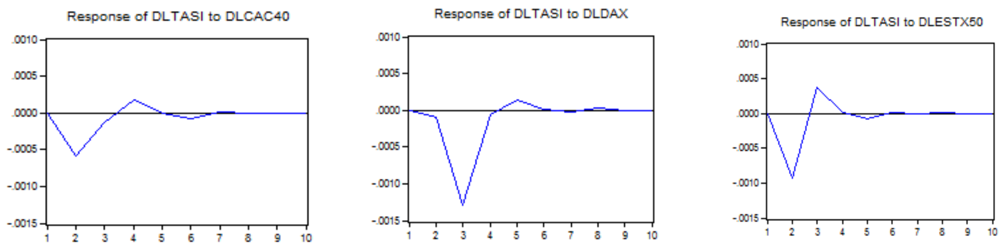

Fig. 5. Impulse response plots after VaR estimation

The impulse response plots show that a shock on CAC40 affects the Saudi index negatively by approximately 0.071% immediately after one time period. Whereas, the impact of DAX on TASI is about -0.017% after two time periods. The impulse response plots also show that a shock on the Euro stocks50 affects the Saudi index by approximately 0.019% immediately and negatively. It is important to note that the study of the impulse response functions for the other European indices shows minor effects on the Saudi index.

4.2.3. Variance decomposition

To complete this empirical analysis of the crisis impact on the Saudi financial market, we propose a variance decomposition after vector autoregressive (VAR) model estimation. This approach is used to analyze the contribution of each European index in the variance of TASI after, during, and before the European crisis period. The variables are considered after a log-difference transformation.

Table 7. Variance decomposition of TASI before European crisis.

Period	S.E.	TASI	ESTX 50	CAC 40	DAX	FTSE _MIB	FTSE _ASE	ISE	PS I20	IBEX 35
1	0.0197	100.000	0.0000	0.0000	0.0000	0.0000	0.0000	0.0000	0.0000	0.0000
2	0.0200	99.0199	0.2127	0.0858	0.0023	0.1059	0.3774	0.0346	0.0974	0.0640
3	0.0201	98.4389	0.2460	0.0901	0.4162	0.1448	0.3766	0.1246	0.0989	0.0640
4	0.0201	98.3355	0.2458	0.0969	0.4164	0.1472	0.3767	0.1296	0.0990	0.1529
5	0.0201	98.2796	0.2474	0.0969	0.4212	0.1475	0.3797	0.1315	0.0990	0.1972
6	0.0201	98.2501	0.2474	0.0982	0.4211	0.1475	0.3820	0.1572	0.0994	0.1972
7	0.0201	98.2299	0.2473	0.0982	0.4213	0.1481	0.3821	0.1699	0.0994	0.2039
8	0.0201	98.2281	0.2474	0.0982	0.4214	0.1481	0.3832	0.1699	0.0994	0.2043
9	0.0201	98.2256	0.2474	0.0982	0.4215	0.1483	0.3834	0.1720	0.0994	0.2043
10	0.0201	98.2255	0.2474	0.0982	0.4215	0.1483	0.3834	0.1720	0.0994	0.2043

Table 8. Variance decomposition of TASI during European crisis.

Period	S.E.	TASI	ESTX50	CAC40	DAX	FTSE _MIB	FTSE _ASE	ISE	PSI20	IBEX35
1	0.0158	100	0	0	0	0	0	0	0	0
2	0.0162	97.8864	0.5586	0.1256	0.3035	0.0489	0.0842	0.2168	0.0042	0.7719
3	0.0165	96.2186	1.1386	0.1670	0.3364	0.1369	0.1017	0.4697	0.3701	1.0610
4	0.0165	95.8088	1.2025	0.2061	0.4434	0.1818	0.1645	0.5206	0.3697	1.1025
5	0.0166	95.5553	1.2642	0.2056	0.4892	0.1866	0.1665	0.6599	0.3691	1.1037
6	0.0166	95.4687	1.2835	0.2087	0.5312	0.1963	0.1665	0.6594	0.3770	1.1088
7	0.0166	95.4321	1.2863	0.2089	0.5315	0.1976	0.1681	0.6848	0.3810	1.1097
8	0.0166	95.4251	1.2872	0.2102	0.5363	0.1977	0.1681	0.6848	0.3811	1.1096
9	0.0166	95.4220	1.2878	0.2103	0.5363	0.1978	0.1685	0.6866	0.3811	1.1096
10	0.0166	95.4214	1.2879	0.2106	0.5364	0.1978	0.1685	0.6867	0.3811	1.1096

Table 9. Variance decomposition of TASI after European crisis.

Period	S.E.	TASI	ESTX50	CAC40	DAX	FTSE _MIB	FTSE _ASE	ISE	PSI20	IBEX35
1	0.0146	100	0	0	0	0	0	0	0	0
2	0.0147	99.3799	0.0022	0.1870	0.2123	0.0798	0.0319	0.0183	0.0332	0.0555
3	0.0147	98.7628	0.0025	0.1859	0.4389	0.3834	0.0582	0.0507	0.0395	0.0779
4	0.0147	98.7248	0.0030	0.2015	0.4489	0.3869	0.0611	0.0521	0.0395	0.0822
5	0.0148	98.6904	0.0031	0.2182	0.4552	0.3870	0.0636	0.0558	0.0396	0.0871
6	0.0148	98.6878	0.0032	0.2181	0.4552	0.3873	0.0636	0.0578	0.0396	0.0873
7	0.0148	98.6856	0.0032	0.2181	0.4568	0.3875	0.0639	0.0578	0.0397	0.0874
8	0.0148	98.6851	0.0032	0.2182	0.4571	0.3875	0.0639	0.0578	0.0397	0.0875
9	0.0148	98.6850	0.0032	0.2182	0.4571	0.3875	0.0639	0.0578	0.0397	0.0875
10	0.0148	98.6850	0.0032	0.2182	0.4571	0.3875	0.0639	0.0578	0.0397	0.0876

Notes: Values in tables provide information about the relative importance of each random innovation in affecting the variation of the variables in the VAR. The variance decomposition is made over ten periods.

Tables 7 to 9 provide the results of the variance decomposition of the TASI after, during, and before the European crisis, respectively. The results show that the variance is changing from one period to another. We observe that the variance of TASI increases during the crisis period. This observation can be explained by the fact that the uncertainty decreases during the crisis and, therefore, the volatility decreases as well. Additionally, the results show that the contribution of the European indices returns in the TASI volatility is more significant during the period crisis than during the period before the crisis. Particularly, the contribution of the Euro-Stoxx index return in the volatility of the Saudi index passes from 0.212% before the crisis to 0.558% during the crisis period. This variation is important and suggests the presence of the European crisis impact on the Saudi financial market.

To summarize, the European crisis undoubtedly affects the Saudi financial market and, as one of the most prominent markets in the world, the Saudi stock market has fallen. Evidently, each financial crisis is followed by an economic crisis, particularly for those economies that are strongly integrated into the global economy, such as the Saudi Arabia economy. This integration may expose Saudi economy to an adverse impact from the world and, especially, from Europe.

In the following section we try to investigate the adverse effects of the European crisis on the Saudi macroeconomic variables.

4.2.4. Dependence and long run effects

To investigate the co-movements of the European and Saudi indices and the long run effects of permanent shocks, first we apply the test proposed by Stock and Watson (1998), and we subsequently follow the methodology implemented by Warne (1993) and based on the Stock and Watson common trends test.

Various bivariate models were estimated and Stock Watson's test was performed to detect possible co-movements between the European and Saudi indices. As indicated in the Table 10, Stock and Watson's test shows two common trends that reflect the presence of a strong dependence between the Saudi and various European financial markets. These results motivate the investigation of the long run effects of permanent shocks.

Table10. The stock Watson test for presence of common trends.

Dependent	Statistic test	Lags	NCT
TASI-CAC40	-2229.049	3	2
TASI-DAX	-2229.058	3	2
TASI-FTSE_MIB	-2229.061	3	2
TASI-FTSE_ASE	-2229.063	3	2
TASI-ISE	-2229.061	3	2
TASI-PSI20	-2229.058	3	2
TASI-IBEX35	-2229.060	3	2
TASI-ESTX50	-2229.064	3	2

[*] As mentioned in the original paper of Stock and Watson (1998), the significance level for two common trends at 1%, 2.5% are respectively -13.8, -10.6. Dependent: corresponds to dependent variables used in the estimated model. The variable Lags corresponds to number of lags used in Prewhitening Filter and NCT is the number of common trends.

The estimation of the common trends model is performed following the methodology proposed by Warne (1993), which is usually used to estimate parameters in common stochastic trends for analysing the long-run effects of permanent shocks. The main results are presented in the Table 11, where the estimated elements of the long run impact matrix are reported. ψ_1 and ψ_2 are assumed to be permanent shocks.

Table11. Long- run effects of permanent shocks.

Variables	ψ_1	ψ_2
LTASI	0.00216(0.0039)	0.9978(0.0039)
LESTX50	0.145(0.058)	0.8543(0.0583)
LCAC40	0.892(0.018)	0.1071(0.0181)
LDAX	0.9995(7.25e^{-05})	0.000406(7.25e^{-05})
LFTSE_MIB	0.999(3.254e^{-05})	0.000182(3.254e^{-05})
LFTSE_ASE	0.187(0.166)	0.812(0.166)
LISE	0.1272(0.147)	0.8727(0.147)
LPSI20	1.000(0.000)	0.000(0.000)
LIBEX35	0.9996(0.000155)	0.00032(0.000155)

Note: * denotes statistically significant at the 5% level.
Asymptotic standard errors are in parentheses.

A number of interesting features stand out after this analysis. For the Saudi index, the long-run variance is explained by only 0.2% of the European crisis (permanent shock ψ_1). However, the permanent shock ψ_2 explains 99.97% of the long-run variance. In contrast, for the European indices in periods when permanent shocks have significant effects (all series except FTSE_ASE and ISE), the results show that the shock induced by the European crisis has a stronger effect than that by other unspecified events. For example, the crisis shock explains the long-run variance of DAX by 99.95%.

4.3. The propagation of the financial crisis to the real economy

- Identifying the transmission channel

To assess the transmission of the financial crisis to the real economy, we primarily examine the impact of the crisis on Saudi Arabia output, inflation and the interest rates. The economic literature has associated the three transmission channels of the financial crisis to the output. The first channel is through the exclusion from international capital markets. The second way is through an increase in the borrowing costs. The last one is the international trade. The results of various empirical studies on the impact of the financial crisis support a significant contraction in output.

In this context, to study the transmission of the financial crisis on the macroeconomic variables (output, export, and import), we apply a non-stationary test in the presence of structural break to detect the crisis period (see the results in Table 3). Next, we estimate a model with a dummy variable to identify the transmission channel of the financial crisis to the real economy. The estimated model is written as

$$y_t - \sum_{t=1}^{T} y_{t-i} = \alpha \ + \beta D_t + \delta X_t + \varepsilon_t .$$

In this specification y_t is the log of real GDP for Saudi Arabia at time *t*, D_t is a dummy variable that takes value 1 before the crisis and 0 otherwise for the first specification, and 1 during crisis and 0 otherwise for the second specification. X_t is a set of exogenous variables that may influence the short term GDP growth. In this modeling the variables introduced in the vector X are trade openness measured as the ratio of total export to import, real exchange rate, interest rate and the lags of GDP. We retain only two lags for endogenous.

As mentioned above, co-integration tests show the presence of structural break in the export and import series (Tables 1 and 3). This result implies that the European crisis does significantly affect the export and import operations between Saudi Arabia and the European countries. In fact, the European Union can be considered as the largest trading partner (the third destination for exports and the first source of imports). Particularly, in the crisis period the Saudi Arabia recorded a trade deficit with the European market. This result can be argued by slower growth in the European counties that reduces the demand for oil and non-oil export from Saudi Arabia. Moreover, these tests show the presence of a breakpoint in the Saudi GDP series revealing the presence of a possible structural change due to the crisis. Additionally, the presence of this structural change can be regarded as a consequence of transmission of the financial crisis on the Saudi output.

To test these hypotheses, we estimate two models in the presence of dummy variables. In the first model, we assume presence of two periods, before and after the crisis. In the second model, the periods are different with a dummy variable: equal to 1 if *t* is during the crisis and 0 otherwise. The modeling results are as follows:

Table 12. Estimation Results for the two dummy variable models.

Variables	**Model1:** $dt = \begin{cases} 1 & if\ \ t\ before\ crisis \\ 0 & otherwise \end{cases}$	**Model 2:** $dt = \begin{cases} 1 & if\ \ t\ during\ crisis \\ 0 & otherwise \end{cases}$
C	-3.696 (0.516)***	-3.113 (0.740)***
Dlgdp_KSA (-1)	-0.277 (0.126)***	-0.190(0.154)
Dlgdp _KSA(-2)	-0.650 (0.114)***	-0.619 (0.143)***
Dlgdp _KSA(-3)	-0.161(0.119)	-0.046(0.144)
DLGDP_KSA(-4)	-0.262 (0.107)***	-0.242(0.140)
dt	-0.250 (0.076)***	-0.017 (0.033)***
Open	4.312 (0.534)***	3.831 (0.730)***
USD_EURO	-0.369(0.276)	-0.902 (0.303)***
TB3M_KSA	-0.042 (0.020)***	0.018 (0.008)***
R-squared	0.850	0.764
Adjusted R-squared	0.783	0.659
F-stat	12.727(0.000)	7.279(0.000)

Notes: ***, **, and * denote significance level at 1%, 5% and 10%, respectively. The values in parentheses denote the Standard errors.

One of the results which stands out most prominently is that the estimated parameter associated with the dummy variable is statistically significant in the two specifications. This finding supports the effect of the crisis on the output. This result suggests that the financial crisis significantly reduces contemporaneous output growth by approximately 2.5%. The first model also shows that the output is significantly determined endogenously because the parameters associated with the first, second, and fourth lags are significant at 1% significance level. In addition to this finding, the results illustrate a significance of the estimated parameter associated the openness variable suggesting that the primary transmission channel of the financial crisis on the Saudi output is international trade. In fact, the main reason of the decreasing in the Kingdom's growth economy is due to the deterioration of export to the European countries in the crisis period. This result is confirmed by the Zivot Andrews test that allows for the presence of structural break in both export and import operations.

The decline in Saudi export can be explained by a direct impact of a significant economic slowdown in the European countries during the crisis and therefore a decline in demand for oil. Moreover, as depicted in the Fig. 3, the oil price has sharply dropped during the crisis period causing shrinkage in the volume of export. To be more precise, oil prices plunged to just over $30 per barrel in the fourth quarter of 2008 in the period immediately following the worst of the financial crisis.

- The impact of the financial crisis on inflation

To investigate the crisis impacts on inflation, we proceed by an estimation of the stylized Phillips curve, defined as a direct relationship between the change of the inflation rate and growth in real output. We also introduce a dummy variable in the model to study the effect of the crisis on the Phillips relationship. The specification used is written as: $d\pi_t = \alpha + \varphi_i \sum_{i=1}^{T} dy_{t-i} + \delta D_t + \nu_t$ where $d\pi_t$ is the log consumer price after second differentiation for Saudi Arabia, dy_{t-i} denote lagged first differenced log GDP, and D_t is a dummy variable that takes value 0 before the crisis and 1 otherwise. To identify the short term crisis impact we retain two lags.

The study consists of testing the erosion of potential output after the European crisis. Under normal economic conditions, faster growth has an immediate consequence on acceleration of inflation, thus the predicted signs of the lag coefficient φ_i are positive. However, if the coefficient δ is statistically significant, we support a crisis impact on inflation rates in Saudi Arabia.

To insure that the estimated regression is not fallacious we investigate the long run relationship between output and inflation by implementing co-integration test.

Table 13. Short run versus long run estimated Phillips curve.

Model 1. Short run			**Model 2.** Long run (Co-integration test)		
Variables	Coeff	Std	Hypothesis	Trace statistic	Prob
DLGDP_KSA(-1)	0.868	0.392	None*	20.925	0.0069
DLGDP_KSA(-2)	0.328	0.391	At most 1*	9.296	0.0023
DT	-0.069	0.047			
Adjusted R-squared	0.225				

Note: * imply rejection of the hypothesis at 5% significance level.

As shown in table 13, both trace and Eigen value statistics highlight 2 co-integrating equations at 5% level. Thus, there exist a long run relationship between the output and inflation. Otherwise, the estimation of the stylized Phillips curve is not fallacious and then we can proceed by a short-run regression.

The short-run regression results show that the estimated coefficient δ associated to the dummy variable is not statistically significant supporting that the crisis does not affect the structure of the relationship between the inflation and the output. Otherwise, the stylized Phillips curve remains unchanged after crisis. This result does not mean that the crisis did not affect Saudi inflation levels. However, it significantly rejects a shift in the estimated regression. The results show also that the sign of the first lag coefficient associated to the output is statistically significant indicating that a faster growth of output causes an acceleration of inflation.

5. Conclusions

In this paper, we used the Gregory and Hansen co-integration test with the presence of structural breaks and the test for common trends proposed by Stock and Watson (1988) to investigate the contagion effect of the European debt crisis on the Saudi stock market. We also estimated a model with a dummy variable to identify the transmission channel of the financial crisis to the Saudi Arabia real economy.

The analysis of the Gregory and Hansen co-integration test results provided substantial evidence in favor of the financial contagion of the Eurozone crisis to the Saudi Arabian markets. Our results highlight that for the majority of the financial series, the breakpoint is detected in May of 2008 corroborating those obtained in Kolb (2011) and Thao and Daly (2012). These researchers detected the breakpoint date in October of 2009. Moreover, the co-integration results with the presence of structural break reveals an apparent impact of the European crisis on the Saudi financial market.

Stock and Watson's test revealed two common trends reflecting the presence of strong dependence between the Saudi and various European financial markets. More specifically, the last test that allows for examination of the long-run variance has illustrated that the shock induced by the European crisis has a stronger effect than other unspecified events.

Another interesting result of this study is that the primary transmission channel of the financial crisis on the Saudi output was the international trade. This result can be supported by the fact that the main reason for the decrease in the Kingdom's growth economy is due to the deterioration of export to the European countries during the crisis period. This result is also confirmed by the Zivot and Andrews test that enables the presence of structural break in both export and import operations. The decline in Saudi export can be explained by a direct impact of a significant economic slowdown in the European countries during the crisis and therefore a decline in demand for oil.

Finally, in this study, the estimated stylized Phillips curve remains unchanged after the crisis, which indicates that the crisis does not affect the structure of the relationship between the inflation and the output.

Our results are of particular importance today, but the study has to be continued because the Eurozone crisis is not over; therefore, the impact on the countries over the world and, especially, on the kingdom may increase or decrease. This impact depends on the economic power of the country, as well as the economic policies proposed by the government, to reduce the long-term effects of the crisis.

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Annex

Table 4
Results of Gregory and Hansen co-integration test.

Independents		**LTASI (dependent)**	
	ADF^*	$Z_\alpha(\tau)$	$Z_t(\tau)$
Model 1. Level shift ***(C)***			
EURO STOXX 50	-3.908	-26.385	-4.084
CAC 40	-4.217	-29.335	-4.237
DAX	-4.043	-32.211	-4.613
FTSE MIB	-3.691	-22.537	-3.770
FTSE/ASE	-3.545	-21.706	-3.826
ISE	-4.076	-25.064	-4.103
PSI-20	-3.618	-23.078	-3.889
IBEX 35	-3.623	-23.158	-3.874
Model 2. Level shift with trend ***(C/T)***			
EURO STOXX 50	-4.023	-29.634	-4.336
CAC 40	-4.307	-31.398	-4.412
DAX	-4.236	-32.368	-4.516
FTSE MIB	-4.180	-30.372	-4.394
FTSE/ASE	-3.753	-26.625	-4.164
ISE	-4.151	-26.730	-4.237
PSI-20	-3.716	-25.495	-4.060
IBEX 35	-3.665	-25.110	-3.984
Model 3. Regime shift ***(C/S)***			
EURO STOXX 50	-4.419	-36.778	-4.419
CAC40	-4.535	-39.920	-4.560
DAX	-4.503	-35.00	-4.660
FTSE MIB	-4.033	-30.069	-4.243
FTSE/ASE	-3.718	-28.831	-4.342
ISE	-4.627	-34.680	-4.323
PSI-20	-3.968	-26.220	-3.921
IBEX 35	-4.042	-31.846	-4.145

Note: Dependent and Independent variables are considered after logarithmic transformation. The critical values for ADF* and z_t^* statistic tests for the three specifications (C, C/T and C/S) as indicated by Gregory and Hansen are -5.44, -5.80, -5.97, -4.92, -529 , -5.50, at 1%, 5% level respectively. The critical values for z_α^* statistic tests for the three specifications (C, C/T and C/S) are -57.01, -64.77,-68.21, -46.98,-53.92 and -58.33 at 1%, 5% level respectively.

Table 5
Results of Gregory and Hansen co-integration test.

Independent	**TASI (Dependent)**		
	ADF^*	$Z_\alpha(\tau)$	$Z_t(\tau)$
Model 1. Level shift ***(C)***			
EURO STOXX 50	-23.029	-2281.66	-48.556
CAC 40	-18.232	-1977.037	-42.234
DAX	-18.245	-1978.150	-42.255
FTSE MIB	-18.233	-1977.710	-42.247
FTSE/ASE	-18.239	-1977.444	-42.242
ISE	-18.238	-1976.881	-42.231
PSI-20	-18.248	-1979.381	-42.279
IBEX 35	-18.243	-1975.974	-42.213
Model 2. Level shift with trend ***(C/T)***			
EURO STOXX 50	-18.406	-1983.034	-42.348
CAC 40	-18.384	-1983.312	-42.354
DAX	-18.398	-1984.399	-42.375
FTSE MIB	-18.386	-1983.898	-42.366
FTSE/ASE	-18.391	-1983.551	-42.359
ISE	-18.394	-1983.145	-42.351
PSI-20	-18.401	-1985.514	-42.397
IBEX 35	-18.395	-1982.056	-42.330
Model 3. Regime shift ***(C/S)***			
EURO STOXX 50	-18.257	-1976.103	-42.215
CAC 40	-18.236	-1996.965	-42.232
DAX	-18.247	-1977.959	-42.252
FTSE MIB	-18.233	-1977.771	-42.248
FTSE/ASE	-18.243	-1977.553	-42.244
ISE	-18.237	-1977.349	-42.240
PSI-20	-18.248	-1979.466	-42.281
IBEX 35	-18.252	-1976.532	-42.224

Note: Dependent and independent variables considered after logarithmic difference; variables in this case correspond to the index returns.

Table 6
Results of Gregory and Hansen co-integration test.

Dependent	**TASI(Independent)**		
	ADF^*	$Z_\alpha(\tau)$	$Z_t(\tau)$
*Model 1. Level shift**(C)***			
EURO STOXX 50	-23.107	-2285.182	-48.634
CAC40	-23.289	-2313.706	-49.213
DAX	-22.399	-2208.346	-46.899
FTSE MIB	-22.454	-2208.748	-46.947
FTSE/ASE	-33.651	-2082.440	-44.285
ISE	-20.876	-2130.574	-45.329
PSI-20	-17.856	-2085.222	-44.379
IBEX 35	-23.115	-2170.917	-46.128
Model 2. Level shift with trend ***(C/T)***			
EURO STOXX 50	-23.342	-2292.617	-48.797
CAC 40	-23.467	-2319.767	-49.347
DAX	-22.620	-2216.430	-47.070
FTSE MIB	-22.734	-2218.582	-47.151
FTSE/ASE	-33.753	-2086.987	-44.373
ISE	-21.197	-2139.532	-45.515
PSI-20	-17.967	-2088.196	-44.440
IBEX 35	-23.226	-2175.015	-46.214
Model 3. Regime shift ***(C/S)***			
EURO STOXX 50	-23.110	-2286.060	-48.653
CAC40	-23.315	-2314.64	-49.235
DAX	-22.408	-2208.720	-46.904
FTSE MIB	-22.569	-2215.334	-47.087
FTSE/ASE	-33.656	-2082.015	-44.277
ISE	-20.881	-2131.349	-45.345
PSI-20	-17.866	-2086.424	-44.404
IBEX 35	-23.112	-2173.624	-46.183

Note: Dependent and independent variables considered after logarithmic difference; variables in this case correspond to the index returns.